Credit Market Disequilibrium in Poland: Can We Find What We Expect? Non-Stationarity and the Short-side Rule*

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Abstract

This paper presents an empirical investigation of the disequilibrium hypothesis on the Polish loan market in the 1990s. Using data over this period of rapid and sustained transition, we estimate a disequilibrium model with a standard maximum likelihood method. However, the estimates are highly counter-intuitive as regards the timing of the identified regimes. We show that the gap between the econometric evidence and the expected results may stem from the phenomenon of stochastic non-stationarity in a disequilibrium setting based on the short-side rule. We find that the omission of one non-stationary variable of the cointegrating space or the absence of a "structural" cointegrating relationship in one or both regimes lead to a spurious configuration. In such a case, the wrong use of the standard likelihood function, derived under the stationarity assumption, may lead to non-convergent estimates of structural parameters and, hence, to an erroneous identification of the regimes. Therefore, as a first approach to this problem, we estimate a disequilibrium model with stationary data, and identify the disequilibrium as an excess of quantities of new loans supplied (or demanded) on the market at time t. The empirical results are then robust and economically founded and correspond to the set and the timing of expected regimes.

- *Keywords* : disequilibrium, monetary standard and regimes, non-stationarity and cointegration, transition, Poland.
- *J.E.L Classification* : D50, E42, C32, P00

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1 Introduction

The Polish credit market unquestionably experienced sharply contrasting developments in the 1990s. The difficulty of the monetary authority's control over credit activity was a permanent concern during the Polish transition process, but the reasons for this changed dramatically over time. Between 1994 and 1998, the combination of fixed exchange rate policies and strong capital inflows led to a spectacular increase in the foreign reserves of the central bank. The amount of gross official reserves expanded at an average annual rate of over 52 per cent. These liquidity inflows triggered a boom in the credit market. The total amount of loans skyrocketed with, on average, a 20 per cent annual increase. At the same time, the steady improvement in the results and health of banks along with harsh competition for customers had a highly favorable impact on changes in loan supply. The loan supply growth rate was so sensitive that it was largely disproportionate as compared with changes in loan demand over the period under review. Between 1995 and 1998, the banks' main problem was to find a reliable, creditworthy borrower rather than the necessary funds for its financing. However, this situation changed suddenly in 1999. First, there was an increase in the instability of real activity. While the average annual growth rate of GDP was nearly 6 per cent in the 1994-1998 period, it decreased by over 30 per cent, to 4.1 per cent in 1999 and 4.0 per cent in 2000. It fell to only 1 per cent in 2001, reaching an all-time low of 0.2 per cent in the fourth quarter. Second, those economic events were concomitant with a tremendous deterioration in the supply side of the credit market. Indeed, the ratio of non-performing loans to total loans rose by almost 26 per cent between 1998 and 1999 and its average annual increase was about 19 per cent over the 1999-2001 period. Moreover, banks' profitability, as measured by the gross profit to income ratio tumbled by nearly 70 per cent between the 1995-1998 and 1999-2001 reference periods. On the whole, since the end of 1999 banks have tried to underpin their activity by contracting their credit supply and investing in risk-free assets. For these different reasons, it is highly probable that the Polish credit market was characterized by a disequilibrium state in the 1990s. The symptoms are so obvious that this is a textbook case for applying the disequilibrium theories developed in the 1970s.

The first research objective of this paper is to confirm econometrically the switch observed in the Polish loan market and to test the disequilibrium hypothesis. In this respect, we apply standard methods derived from a behavior rule, the short-side rule, based on the voluntary exchange principle. The underlying idea is that the observed amount of loans is equal to the minimum of the quantities demanded and supplied, which are assumed to be unobservable. A substantial strand of the literature has been devoted to the estimation of these sorts of models since the 1970s. Here, we adopt the standard maximum likelihood estimation method developed by Maddala and Nelson (1974). Yet, we show that the implementation of this technique does not enable us to identify the type of disequilibrium *a priori* expected on the Polish loan market. The results are highly counter-intuitive, in particular as far as the identification of the regimes in the February 1994 - February 2002 period is concerned. Indeed, for all different specifications we obtain the opposite regimes to those expected. The model generates a high estimated probability of a supply regime over most of the period under consideration and a demand regime at the end of the period. These findings are robust to the specification choice but are completely at odds with the history of the Polish loan market over the sample period. Therefore, it is important to understand the reasons for the discrepancy between the econometric evidence and the expected results.

The main aim of our paper is to give an insight into this empirical puzzle. More precisely, draw attention to the phenomenon of stochastic non-stationarity in a disequilibrium setting. In fact, it appears that almost all of our series are non-stationary. In addition, the assumption of cointegration is strongly rejected for most of the equilibrium linear specifications studied. Consequently, the question is to know how to use the concept of cointegration within a non-linear configuration such as the short-side rule given that, by definition, the loan demand and supply variables are unobservable.

In fact, we think that it is highly probable that there is no cointegrating relationship between the unobservable variables and the set of explanatory variables specified in each regime. Yet, in that case (stemming from the absence of a "structural" cointegrating relationship or from the omission of one non-stationary variable of the cointegrating space), it is no longer possible to estimate the disequilibrium model using the maximum likelihood methodology proposed by Maddala and Nelson (1974). When residuals in at least one regime are non-stationary, the log-likelihood function is then asymptotically degenerated and its use in a estimation procedure would lead to a spurious estimation. As a result, the identified demand and supply regimes, based on non-convergent estimates of structural parameters, may be aberrant. Ultimately, what is at stake is that one cannot test *ex-post* or *ex-ante* for the existence of a cointegrating relationship in each regime, i.e. for the stationarity of the residuals because the "true" demand and supply functions are by nature unobservable.

In this paper, we propose a first approach to dealing with this problem that consists in estimating a disequilibrium model with stationary data (as would be the case in a linear model). The logic is then different from that considered in a standard disequilibrium model. More precisely, given the non-stationarity issue, we propose to identify the disequilibrium as an excess of quantities of new loans supplied (or demanded) on the market. For instance, if the growth of supply is lower than that of demand, changes in the observed quantities of loans are assumed to be equal to the less restrictive growth rate, i.e. the growth of the quantities supplied, irrespective of the level of loan supply and demand. Naturally, this definition does not correspond to the standard definition of disequilibrium based on the voluntary exchange principle. We discuss the similarities and divergences between both approaches. However, we observe that our approach has many advantages. First, it appears that the results obtained are robust and economically founded. In the most sophisticated specifications all coefficients have the expected sign and virtually all variables are statistically significant. Second, and more importantly, the set and the timing of regimes we identify are in line with the stylized facts pertaining to the history of the Polish loan market. Even though this solution to the issue of non-stationarity in the absence of a cointegrating relationship (spurious regression configuration) is not completely satisfactory, it is a first step in the right direction. It clearly shows that in the Polish case, if relevant, the regimes must be identified with the stationary components of both the amount of credit and the observable explanatory variables, and not directly with the variables in levels.

The paper is organized as follows. In Section 2, we briefly recall the traditional maximum likelihood method for estimating models of markets in disequilibrium. In Section 3, the results obtained on the Polish credit market are presented and compared to the main stylized facts. Section 4 describes the main aspects of the non-stationarity in a disequilibrium representation. In Section 5, we propose estimates for a disequilibrium model with stationary data and Section 6 concludes the paper.

2 Disequilibrium econometrics: the short-side rule

Since Fair and Jaffee (1972) a large body of literature has been devoted to the econometric problems associated with estimating demand and supply schedules in disequilibrium markets. The main approach consists in using some maximum likelihood (ML) methods. In a seminal paper, Maddala and Nelson (1974) derived the general likelihood function for different disequilibrium models and proposed the appropriate ML estimating procedures. The simplest model considered by the authors is as follows:

$$d_t = x'_{1,t}\beta_1 + \varepsilon_{1,t} \tag{1}$$

$$s_t = x'_{2,t}\beta_2 + \varepsilon_{2,t} \tag{2}$$

$$q_t = \min(d_t, s_t) \tag{3}$$

where d_t denotes the unobservable quantity demanded during period t, s_t the unobservable quantity supplied during period t, $x'_{1,t} = \left(x^{(1)}_{1,t}\ x^{(1)}_{2,t} ... x^{(1)}_{K_1,t}\right)'$ is a vector of K_1 explanatory variables that influence d_t, $x'_{2,t} = \left(x^{(2)}_{1,t}\ x^{(2)}_{2,t} ... x^{(2)}_{K_2,t}\right)'$ is a vector of K_2 explanatory variables that influence s_t, β_1 and β_2 are respectively $(K_1, 1)$ and $(K_2, 1)$ vectors of parameters. We assume that d_t and s_t are unobservable at date t, whereas $x_{1,t}$ and $x_{2,t}$ are observable. The variable q_t denotes the actual quantity observed at time t. The equation (3) is the crucial disequilibrium hypothesis, which allows for the possibility that the price of the exchanged good is not perfectly flexible and rationing occurs. More generally, equation (3) indicates that any disequilibrium which takes place, i.e. any divergence between the quantity supplied and demanded, results from lack of complete price adjustment. Therefore, on the basis of voluntary exchange the "short side" of the market prevails.

Because of the equation (3), the model itself determines the probabilities with which each observation belongs to either supplied or demanded quantities. Following this, we briefly develop the theoretical underpinnings of this result. In a first version of the model, Maddala and Nelson (1974) assume that both residuals $\varepsilon_{1,t}$ and $\varepsilon_{2,t}$ are stationary processes, independently and normally distributed with variance σ_1^2 and σ_2^2 respectively. Under these regularity assumptions, the transformed variable $\varepsilon_{1,t} - \varepsilon_{2,t}$ is normally distributed with a variance equal to $\sigma^2 = \sigma_1^2 + \sigma_2^2$. Hence, the reduced variable $(\varepsilon_{1,t} - \varepsilon_{2,t})/\sigma$ follows a $N(0,1)$. Then, the probability that the observation q_t belongs to the demand regime, denoted $\pi_t^{(d)}$, can be computed as the corresponding $N(0,1)$ fractile:

$$\pi_t^{(d)} = P(D_t < S_t) = \Phi(h_t) = \frac{1}{\sqrt{2\pi}} \int_{-\infty}^{h_t} e^{-\frac{x^2}{2}} dx \tag{4}$$

where $h_t = \left(x'_{2,t}\beta_2 - x'_{1,t}\beta_1\right)/\sigma$, and $\Phi(x)$ denotes the cumulative distribution function of the $N(0,1)$. Symmetrically, the probability of obtaining the supply regime, denoted $\pi_t^{(s)}$, is defined by $P(S_t < D_t) = 1 - \Phi(h_t)$.

Let θ denote the vector of structural parameters $\theta = (\beta_1\, \beta_2\, \sigma_1\, \sigma_2)'$. In order to compute the marginal density, $f_{Q_t}(q_t)$, of the observable variable q_t, we consider the joint density of d_t and s_t, denoted $g_{D_t,S_t}(d_t, s_t)$. Given the definition of the disequilibrium, we know that:

$$f_{Q_t}(q_t) = f_{Q_t|D_t<S_t}(q_t) + f_{Q_t|S_t<D_t}(q_t) \tag{5}$$

Then, we obtain the corresponding marginal density of q_t on the two subsets (cf. Appendix (A.1)):

$$f_{Q_t|D_t<S_t}(q_t) = \int_{q_t=d_t}^{\infty} g_{D_t,S_t}(d_t, z)\, dz \tag{6}$$

$$f_{Q_t|S_t<D_t}(q_t) = \int_{q_t=s_t}^{\infty} g_{D_t,S_t}(z, s_t)\, dz \tag{7}$$

Finally, we obtain the unconditional density function of Q_t:

$$f_{Q_t}(q_t) = f_{Q_t}(q_t, \theta) = \int_{q_t}^{\infty} g_{D_t,S_t}(q_t, z)\, dz + \int_{q_t}^{\infty} g_{D_t,S_t}(z, q_t)\, dz \tag{8}$$

Next, conditionally to a structural parameters set θ and a sample of observable variables $q_t, x_{1,t}$ and $x_{2,t}$ observed on T periods, the log-likelihood function of the model is then defined by:

$$L(\theta) = \sum_{t=1}^{T} \log\left[f_{Q_t}(q_t, \theta)\right] \tag{9}$$

If we assume that both residuals ε_1 and ε_2 are independent, the unconditional density function of Q_t can be expressed as follows:

$$\begin{aligned} f_{Q_t}(q_t) &= \frac{1}{\sigma_1}\phi\left(\frac{x'_{1,t}\beta_1 - q_t}{\sigma_1}\right)\Phi\left(\frac{x'_{2,t}\beta_2 - q_t}{\sigma_2}\right) \\ &\quad + \frac{1}{\sigma_2}\phi\left(\frac{x'_{2,t}\beta_2 - q_t}{\sigma_2}\right)\Phi\left(\frac{x'_{1,t}\beta_1 - q_t}{\sigma_1}\right) \end{aligned} \tag{10}$$

where $\phi(.)$ denotes the normal $N(0,1)$ density function. The proof is provided in Appendix (A.1). In this case, the first and second order derivatives of $L(\theta)$ can be computed analytically (Maddala and Nelson, 1974) or numerically. We can use an iterative procedure such as the Newton-Raphson procedure to obtain the ML estimates of the structural parameters θ. Given the estimated values of the parameters, we can compute the estimated probability that the observation q_t belongs either to the demand or the supply regime, $\widehat{\pi}_t^{(d)}$ and $\widehat{\pi}_t^{(s)}$.

The measurement of credit rationing with the short-side rule and the Maddala and Nelson (1974) estimation technique has been extensively used in the literature. Empirical studies based on aggregate time series data and applied to business loans include Laffont and Garcia (1977) for the Canadian market, Sealey (1979) and King (1986) for the US market. Pazarbaşioğlu (1997) and Kim (1999) investigate whether there was a credit crunch, respectively in Finland following the banking crisis of 1991-92, in Korea following the financial crisis in December 1997. Finally, in a recent paper Atanasova and Wilson (2004) provide panel data estimates of disequilibrium in the UK corporate loan market.

The disequilibrium econometrics approach based on the short-side rule is consistent with the theory of equilibrium credit rationing derived from asymmetric information models. Because of moral hazard and adverse selection problems an excess demand in the loan market may exist, since a rise in the loan interest rate can reduce the expected return of banks. In their seminal paper on credit rationing, Stiglitz and Weiss (1981) also provided a strong intuition why asymmetric information can also lead to persistent excess supply equilibria. The explanation is along the following lines. Let us assume that a bank with an excess supply of loanable funds decreases its interest rate in order to attract new customers, but does not know their creditworthiness. If a competing bank knows who its most profitable customers are, in a countering move it can accept to decrease its interest rate in the same proportion for "good" credit risks, but not for the "bad" ones. Consequently, if each bank can attract only the least profitable customers, each bank may have an excess supply of loanable funds, but no one will accept to reduce its interest rate. All in all, there might be an excess supply of credit in the banking sector, without a fall in the interest rate to clear the loan market.

3 The credit market in Poland and disequilibrium

We now propose to estimate a disequilibrium model on the credit market in Poland during the 1990s.

3.1 Data description and expected signs of the coefficients

Our data set covers the period from February 1994 to February 2002, including 97 monthly observations. All data are obtained from the National Bank of Poland (hereafter NBP), except industrial production ($PROD_t$) and import prices (IMP_t), which come from the Monthly Bulletin of the Polish Central Statistical Office (GUS).

Most of the series used in this study are defined for a set of banks. The sample includes 20 banks in February 1994 and 11 in February 2002 (the difference is due to the consolidation process). It represents on average for the period under consideration, 75.3 per cent of the banking system in terms of total corporate bank debt and 84.7 per cent in terms of total deposits. The variables calculated for our sample of banks pertain to the resident and non-resident sectors and are defined as follows. For the loan series, Q_t, we use total zloty denominated loans up to one year extended to the Polish corporate sector[1]. Among total bank loans granted to the corporate sector, the scope

[1] The "corporate sector" includes state-owned enterprises, private enterprises with up to 9 or more

of the analysis has been restricted to domestic currency loans due to data availability issues (unavailability of interest rates for loans extended in foreign currencies). We will now explain our choice of loan maturities used in this paper. The theoretical underpinnings for the basic loan demand and supply specifications presented here draw on the seminal model of the bank lending channel by Bernanke and Blinder (1988). Yet, in the wake of their paper, empirical tests of this transmission channel based on interest rate spreads mainly consider short-term maturities[2]. However, the Q_t series used in the paper is still highly representative, as it includes, on average for the period under consideration, 36.5 per cent of the total credit extended to firms for our sample of banks, and 27.5 per cent if the whole Polish banking sector is taken into account. The loan interest rate, IL_t, is an arithmetical mean of 3-month, 6-month and 1-year weighted averages for minimum loan rates applied to Polish firms. It is expected to have a negative sign in the demand function and a positive sign in the supply one. As for total deposits, DEP_t, we use the sum of demand and time zloty and foreign currency deposits and of interest on these deposits of both households and firms. This variable is expected to have a positive coefficient in the loan supply equation. Time deposits, SAV_t, are defined only for firms and also include interest on deposits. It is a proxy for firms' cash flow and is expected to have a negative impact on loan demand. More specifically, as SAV_t is composed of liquid assets that can be used in the event of an emergency, it provides an insight into firms' borrowing constraints, everything else being equal. An increase in its value suggests that financial constraints are not binding and that self-financing is a viable option. In contrast, a decrease in its value implies that firms' ability to use internal funds is reduced, thus leading to an increase in their demand for external resources. Lastly, the OL_t variable, concerns only the corporate sector and is defined as the following ratio: zloty denominated loans beyond one year plus total foreign currency loans divided by total loans extended. Hence, it is equal to the part of credit that is excluded from the dependent variable. Its aim is to capture the impact of other loans extended to the corporate sector. Introduced in the loan

employees, cooperatives, and farmers.

[2] Moreover, although our purpose was not to test the existence of the bank lending channel, this paper was written for a PhD of one of the authors. Its aim, *inter alia*, was to discuss the relationship between the bank lending channel and credit rationing. Hence, by considering loans with short-term maturities only, it was possible to compare the different results presented in the dissertation.

demand equation, the corresponding coefficient is expected to display a positive sign in the case of complementary effects or a negative one if other banking means can be substituted for zloty denominated loans up to one year.

The remaining series are defined at the aggregate level. The intervention rate of the central bank, IC_t, is a weighted average of 1 to 14-day reverse repo rates and that of the central bank securities issued for different maturities between February 1994 and January 1998. Since then, it is equal to the actual rate on 28-day NBP bills (also taking into account the average rate of outright sales). Due to the substitution effect, IC_t should have the opposite sign to the loan interest rate, IL_t, in both supply and demand equations. The variable $LNFA_t$ is the ratio of net foreign assets of the central bank to the overall value of banking funds absorbed by sterilization operations conducted by the NBP, which include the net value of reverse repos, outright and NBP bills issue operations and the value of required reserves. It aims to measure on the supply side of the market the net liquidity impact on banks of foreign reserve accumulation by the central bank due to exchange rate interventions. It could be claimed that $LNFA_t$ is no longer an appropriate indicator since the implementation of the floating exchange rate system, i.e. at least June 1999[3]. However, according to OECD (2001), the NBP continued to make off-market foreign exchange transactions as an agent of the government in 2000 and early 2001. This had the same impact on banking liquidity as market interventions. The variable ATB_t measures the share of Treasury bills to total banking assets. It is a rough measure of the imperfect substitutability between bonds and loans mainly due to changes in the riskiness of financial intermediation, and is expected to have a negative coefficient in the loan supply function. With regard to industrial production, $PROD_t$, the theoretical sign of its parameter is indeterminate. In the literature on loan market disequilibrium [4], a lagged index of industrial production is often used to approximate the expectations of firms and the banks about future economic activity and is expected to have a positive coefficient. Following Bernanke and Blinder's (1988) model and bank lending channel literature, a positive dependence of loan demand on output is assumed due to working capital or liquidity considerations. Yet, the latter assumption is rather

[3] Date at which the NBP definitively stopped its foreign exchange market interventions by abolishing fixing transactions.

[4] See, for instance, Sealey (1979), Kim (1999), Pazarbaşioğlu (1997).

ambiguous if corporate sector loan demand is considered. In fact, a drop in industrial production will probably strengthen the liquidity constraint of firms, thus increasing their short-term credit demand. Gertler and Gilchrist (1993), Dale and Haldane (1995) among others, find a short-term increase in bank lending to the corporate sector in the wake of tight money. Finally, the index of import prices, IMP_t, is expected to have a positive coefficient on the demand side of the market. The idea behind introducing this variable is that a change in import prices of intermediate and industrial goods will directly affect the production costs of firms and, hence, their loan demand[5].

In all models the variables are in nominal terms, expressed in logarithm, except for the interest rates which are measured in percentage terms.

3.2 Estimation results

In order to asses the robustness of our results, we consider several specifications of disequilibrium models on the Polish loan market. The results are reported in Table (1).

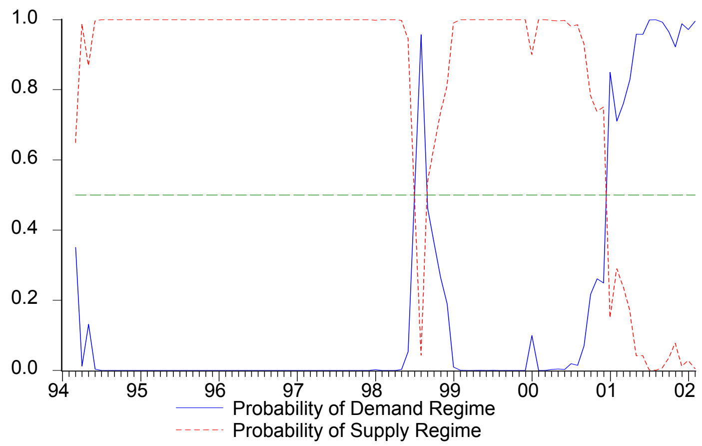
——— Insert Table (1) ———

When applying the Maddala and Nelson (1974) approach to the Polish loan market, the adjusted R-squared statistic is very high (almost equal to 1) and the coefficients of some variables, especially those of interest rates, are not significant and/or do not display the expected sign. Actually, this is particularly true for the loan interest rate in the supply equation. More generally, we find that, irrespective of the side of the market, interest rates do not seem to play a significant role or they may even have an opposite-to-expected impact, which is puzzling[6]. However, these observations are not the most worrying. Our estimates reveal the existence of a succession of two regimes. Figure (1) shows the estimated probabilities of each regime for the best model in terms of the log-likelihood statistics, i.e. for Model 6. It appears that the loan market was

[5]It should be noted that according to Eurostat statistics, the imports of goods and services to GDP ratio in Poland was about 21.5 per cent in 1995 and 34.4 per cent in 2000.

[6]In Model 5, we checked whether these findings are not driven by some collinearity problems between the interest rates. To this end, we used the spread, $IL_t - IC_t$, and the intervention rate, IC_t, instead of IL_t and IC_t in each equation. The results reported in Table (1) are identical to those of Model 4. When introducing a one-lag structure on the interest rates as in Model 6, both interest rates in the loan demand equation have the expected sign and statistically significant coefficients. But, concerning the loan supply interest rate coefficients, the problem still persists.

Figure 1: Estimated Probabilities of Each Regime, III/94 - II/02



characterized by an almost permanent and strong supply regime until the end of 2000. A demand regime has unambiguously emerged since the beginning of 2001. It should be noted that these results seem robust since Models 1, 2, 3, and 4 yielded very similar configurations for both regimes.

These results on the timing of supply and demand regimes are undoubtedly not representative of the recent history of the Polish loan market. First, the fixed exchange rate policies that followed until at least June 1999 can be expected to have had a positive, long-standing and pronounced impact on the loan supply. On the one hand, sterilization of capital inflows has created, since the end of 1993, structural excess liquidity in the banking system, defined as a net indebtedness of the central bank towards commercial banks. They potentially rendered demand regimes much more likely. At the same time, since these capital inflows were in fact imperfectly sterilized[7], they also had a direct favorable impact on the loan supply. In the light of these facts, the finding of a supply regime in the 1990s does not seem plausible.

Second, banks' health has considerably improved since 1994. In fact, the share of non-performing loans in total loans for enterprises and households decreased steadily from 31 per cent in 1993 to 28.7 per cent in 1994 reaching a low of 9.8 per cent in

[7]Bofinger and Wollmershäuser (2002) find in the Polish case an absolute sterilization coefficient equal to 0.89 over the period from December 1991 to September 1998.

September 1997. However, the ratio experienced an upsurge from 10.9 per cent in December 1998 to 13.7 per cent in December 1999, and climbed to 18.3 per cent in December 2001 (for the corporate sector only, the corresponding figures were 11.9, 15.1, and 20.4 per cent, respectively). Manifestly, these stylized facts lead us to expect the opposite regimes to the those obtained by the estimates.

Third, a credit boom occurred in the loan market from mid-1996 to mid-1997. This led the central bank to implement an unprecedented measure (introduction of deposit accounts for the general public) as the tightening of traditional monetary policy instruments did not have the expected impact on banks' lending behavior. Again, these observations strongly suggest that the credit boom was led rather by supply factors and not demand-driven.

——— Insert Table (2) ———

Fourth, although 1993 is often considered in the literature as a typical "credit crunch" year due to the "bad-loans" problem of the Polish banking sector in the wake of the 1990-1991 transformational recession, it does not necessarily imply that this state of the market prevailed in the subsequent years. This lends support to the fact that following the government recapitalization program[8], we should expect the opposite. What needs to be stressed here is that in 1994 the level of credit, measured by the ratio of credit to the corporate sector to GDP (see Table (2)) was extremely low hardly reaching 17.1 per cent of GDP. Therefore, its volume was likely to increase sharply. Moreover, in the same year, there was an almost 40 per cent decline in the ratio as compared with 1989, and a 25 per cent decrease as compared with 1991. As a result, as, since 1994, banks have wanted to build up again their market share on a healthy basis, we should therefore observe a demand rather than a supply regime in the 1990s.

4 Disequilibrium representation and non-stationarity

We must therefore attempt to explain how such unexpected results can be obtained from the standard method of estimating disequilibrium models. There are three main

[8] Taking the average current exchange rate for each year, the goverment recapitalization amouted to USD 1.14 billion in 1993 and USD 1.75 billion in 1994.

competing explanations, which may lie at the root of our striking findings. The first concerns the reliability of the estimation procedure, the second deals with the rationing rule used to represent the credit market disequilibrium and the third concerns the empirical reliability of the disequilibrium assumption for the Polish credit market. However, our interpretation adds a fourth explanation based on the non-stationarity of the data used in this exercise.

4.1 How to interpret the estimation results?

The first source of explanation may be linked to the ML estimation technique used to identify the demand and supply regimes. Two issues are worth mentioning with regard to the estimation technique. The first is linked to the convergence of the iterative procedure. Indeed, the optimization method of the log-likelihood function $L(\theta)$ (equation (9)) should be handled very carefully since the global concavity of $L(\theta)$ is not assured. For instance, it is well known that the global maximum of the likelihood function is infinity if one (or both) of the residual variances converges to zero. Therefore, the application of the standard ML procedure should be adapted in this case as only the local maximum must be searched for. That is why in our application we use three different optimization algorithms in order to evaluate the robustness of our estimates: (i) Newton Raphson with numerical or analytical gradient and Hessian matrix (ii) Nelder-Mead simplex method (iii) Newton Raphson with non-negativity constraints on σ_i.

The second concern is how to determine the starting values used for the optimization algorithms. There are various methods to obtain the initial conditions on structural parameters θ used in the ML optimization. Here, we use a two-step OLS procedure presented in Appendix (A.2). This procedure applied to artificial data leads to the convergence of all the optimization algorithms used and gives estimates, which are very close to the true set of parameters. As a result, this does not seem to be the right direction to find the explanation for our striking empirical results.

The second potential explanation deals with the choice of the short-side rule, which might be an inadequate way to describe the disequilibrium in the loan market. Yet, it seems important to note that this class of models has robust microfoundations, with a theoretical justification based on the voluntary exchange principle. Rather, the problem may stem from the known issue of aggregation. As aggregate data are averages of corporate data, some firms may be credit constrained whereas the average firm may not (Perez, 1998). However, this risk does not seem to be of major relevance in the Polish case. Indeed, monthly surveys of companies performed by the central bank reveal that the share of firms using bank credit is very high: it grew from approximately 80 per cent in 1995 to more than 85 per cent in 1999 (Łyziak, 2001). Moreover, Polish firms are highly dependent on bank credit for their investment activity. In 1998, for instance, 55.2 per cent were entirely dependent on bank financing and 25.9 per cent were able to cover only a smaller part of their investment spending without bank credit. These stylized facts indicate that micro- and macro-level studies on credit rationing would probably yield similar results. As a result, the strong link between real activity and loan market developments described in the introduction of the paper suggests that macro-level evidence on credit rationing must mirror most types of microeconomic behavior, and not be an artificial bias resulting from aggregation problems.

The third concern relates to the empirical validity of the disequilibrium assumption in the Polish case. For the reasons given above, it seems clear that a change in credit market conditions occurred at the end of the 1990s. However, this change does not necessarily mean that the disequilibrium representation matches the Polish stylized facts. Another way to check the disequilibrium hypothesis is to assess the transmission of the monetary policy. If the price of loans is not perfectly flexible and does not clear the loan market, i.e. if the disequilibrium hypothesis is relevant, then we should not observe *a priori* a smooth transmission of monetary policy-controlled interest rates to loan rates. Empirical evidence shows that the responsiveness of loan rates to policy-influenced rates was less than 1 over the period from January 1994 to March 1998 (Opiela, 1999), although the pass-through is found to increase over time as evidenced by Crespo-Cuaresma *et al.* (2004) for a longer data set. Yet, even if both interest rates move in tandem it does not necessarily preclude the existence of a disequilibrium in the credit market. Indeed, the adjustment of the loan rate may be still insufficient in order to clear the loan market.

More generally, this analysis must be put into perspective. In fact, factors other than market disequilibrium may also explain the stickiness of loan interest rates. Lending rates may not move one for one with changes in the intervention rate due to switching costs, i.e. costs implied by changing banks[9]. Implicit risk sharing may be another explanation for the stickiness of interest rates. According to this approach, by offering below-market rates to risk-averse borrowers during periods of tight money, a bank is later compensated when market rates are low with a higher average rate than would be charged to risk-neutral borrowers (Fried and Howitt, 1980). The intensity of competition in the loan market may also play a role: the higher the competition, the smaller the degree of adjustment in the loan rate to changes in the policy rate. Finally, for a given level of imperfect competition, the value of the elasticity of loan demand is another crucial determinant: the greater the availability of substitutes for bank loans, the higher the elasticity of loan demand, the smaller the market power of the bank on loans and, as a consequence, the smaller the sensitivity of the loan rate to changes in the money market rate.

4.2 The non-stationarity hypothesis

Here, we propose another potential explanation for our results. For several reasons, the focus of the discussion is shifted to the issue of non-stationarity of the data used in the estimates. We call attention to this problem since, except for the interest rates, virtually all the observed variables used in the disequilibrium models are $I(1)$ processes. The results of unit root tests are reported in Appendix (A.3). If the loan market was continually in equilibrium, the following standard approach would be adopted. The model would simply be estimated using a linear specification, i.e. by regressing the observed amount of loans on all the explanatory variables of both loan demand and supply functions. The estimates would be made in levels in the presence of a cointegrating relationship and in first differences otherwise. At least in Models 1 and 2, the cointegration tests in linear models indicate the absence of cointegration relationship.

[9] See Klemperer (1995) for an overview of the theoretical literature and Kim *et al.* (2003) for recent empirical evidence on the Norwegian loan market.

So, if the disequilibrium hypothesis is rejected in these models, the linear representation of the observed quantity of loans must be based on the first differences of the non-stationary series. What does it look like in the disequilibrium representation? How can we transpose the concepts of cointegration if the equilibrium assumption is violated and the loan market can be described by a disequilibrium model with a "min" condition as a transition rule? What are the consequences of the use of the standard ML procedure when the cointegration is rejected in one of the two regimes?

In order to assess the influence of non-stationarity in the "short-side" representation, let us consider the initial model (equations (1), (2) and (3)) and assume that all observed explicative variables $x_{j,t}^{(i)}$ are generated by an integrated stochastic process. For simplicity, let us assume a pure random walk process $\forall j = 1, .., K_i, \forall i = 1, 2$:

$$\Delta x_{j,t}^{(i)} = \mu_{j,t}^{(i)} \tag{11}$$

where the innovations processes $\mu_{j,t}^{(i)}$ are *i.i.d.* We also assume that both demand d_t and supply s_t are integrated processes. As a result, two configurations appear: *the cointegrating representation* and *the spurious regression representation.*

In the first case, we have, in each regime, a linear relation between a non-stationary unobservable dependent variable and non-stationary explicative variables $x_{j,t}^{(i)}$. We further assume that, in each regime, there is a cointegrating relationship between these variables represented by:

$$d_t = x'_{1,t}\widetilde{\beta}_1 + \eta_{1,t} \tag{12}$$

$$s_t = x'_{2,t}\widetilde{\beta}_2 + \eta_{2,t} \tag{13}$$

where processes $\eta_{1,t}$ and $\eta_{2,t}$ are $I\left(0\right)$. In this case, the log-likelihood of the observation q_t is well-defined since the regularities assumptions are satisfied. The maximum likelihood methods can be applied to estimate the parameters of the models and to identify the corresponding regimes. If we also assume that there is a cointegrating relationship between the demand d_t, and the supply s_t, i.e. that these unobservable variables share the same common stochastic trend, then the model can be rewritten as:

$$d_t = x'_{1,t}\widetilde{\beta}_1 + \eta_{1,t} \quad (14)$$

$$s_t = d_t + \zeta_t \quad (15)$$

where the process ζ_t is stationary. In this case, we can also apply the Laroque and Salanié (1997) strategy based on a first differences final form.

In the second case, i.e. *the spurious regression representation*, we also have a linear relation between a non-stationary unobservable dependent variable and non-stationary explicative variables $x'^{(i)}_{j,t}$. But we assume that, in each regime, there is no cointegrating relationship between these variables. In other words, we have:

$$d_t = x'_{1,t}\beta_1 + \varepsilon_{1,t} \quad (16)$$

$$s_t = x'_{2,t}\beta_2 + \varepsilon_{2,t} \quad (17)$$

where $\varepsilon_{1,t}$ and $\varepsilon_{2,t}$ are $I(1)$ processes. It should be noted that such a framework does not necessarily imply the absence of a long-run relationship between d_t and s_t. The problem in this case is that the likelihood of the disequilibrium model is asymptotically degenerated since the residuals do not have a stationary distribution. The standard ML procedure is then not appropriate asymptotically since the marginal distribution of q_t (equation (10)) is not well-defined. For a finite sample, when considering wrongly the Maddala and Nelson's likelihood function, the numerical optimization may not converge, or may converge but then leads to non-convergent estimates of the structural parameters. Thus, the identification of the supply and demand regimes becomes impossible or erroneous.

5 The disequilibrium model with stationary data

The previous section illustrated that the non-stationarity of our data could lead, in the absence of a cointegrating relationship between the unobservable variable and the corresponding set of explanatory variables (which is a highly probable outcome), to an erroneous identification of regimes. In order to detect whether such a configuration is responsible for our striking results on the Polish loan market, we propose here to identify the disequilibrium from a regime perspective, with stationary transformations of data (as we would realize in a linear representation). We achieve stationarity of the data by applying long differences (more precisely, annual growth rates) to each series, with the exception of the interest rates, which are stationary in levels. However, in such configuration the main difficulty is to specify the transition rule. More precisely, let us assume that it is necessary (given our non-stationarity hypothesis) to identify the disequilibrium in the loan market with the stationary component of the observed data. There are (at least) two solutions to circumvent the non-stationarity issue in our context.

The first consists in estimating the model defined by the true data generating process (DGP) of the variable Δq_t. If the standard disequilibrium model is valid, the true DGP of the growth rate of the quantity of loans is as follows:

$$\Delta q_t = \begin{cases} \Delta d_t & \text{if } d_t < s_t \text{ and } d_{t-1} < s_{t-1} \\ \Delta s_t & \text{if } s_t < d_t \text{ and } s_{t-1} < d_{t-1} \\ d_t - s_{t-1} & \text{if } d_t < s_t \text{ and } s_{t-1} < d_{t-1} \\ s_t - d_{t-1} & \text{if } s_t < d_t \text{ and } d_{t-1} < s_{t-1} \end{cases} \tag{18}$$

with

$$\Delta d_t = \Delta x'_{1,t}\beta_1 + \mu_{1,t} \tag{19}$$

$$\Delta s_t = \Delta x'_{2,t}\beta_2 + \mu_{2,t} \tag{20}$$

or, equivalently:

$$\Delta q_t = \min(d_t, s_t) - \min(d_{t-1}, s_{t-1}) \tag{21}$$

where Δd_t denotes the annual growth rate of the unobservable demanded quantity of loans during period t, Δs_t the annual growth rate of the unobservable supplied quantity of loans during period t[10]. If the true DGP of q_t is such that $q_t = \min(d_t, s_t)$, the dynamic of any linear transformation of the loan variable follows a four-regime dynamics. Indeed, for the first-difference series Δq_t (or the long difference $\Delta_z q_t = q_t - q_{t-z}$), there are two regimes of status-quo with demand higher (or lower) than

[10] As in section 1, residuals $\mu_{1,t}$ and $\mu_{2,t}$ are assumed stationary and independently and normally distributed.

supply at two points in time (t and $t-1$, or t and $t-z$) and two regimes of switches from demand to supply or vice versa. However, this dynamic model requires a particular method of estimation, since the likelihood is not analytically tractable (Lafont and Montfort, 1979; Laroque and Salanié, 1995). This solution is very interesting and original, but it is technically very hard to implement and appears to be largely beyond the scope of our applied econometric paper.

The second solution, in order to identify the disequilibrium in the loan market and to take into account the non-stationarity of data, consists in using another model, which does not coincide with the traditional disequilibrium model. We propose to use the following model:

$$\Delta d_t = \Delta x_{1,t}' \beta_1 + \mu_{1,t} \tag{22}$$

$$\Delta s_t = \Delta x_{2,t}' \beta_2 + \mu_{2,t} \tag{23}$$

$$\Delta q_t = \min(\Delta d_t, \Delta s_t) \tag{24}$$

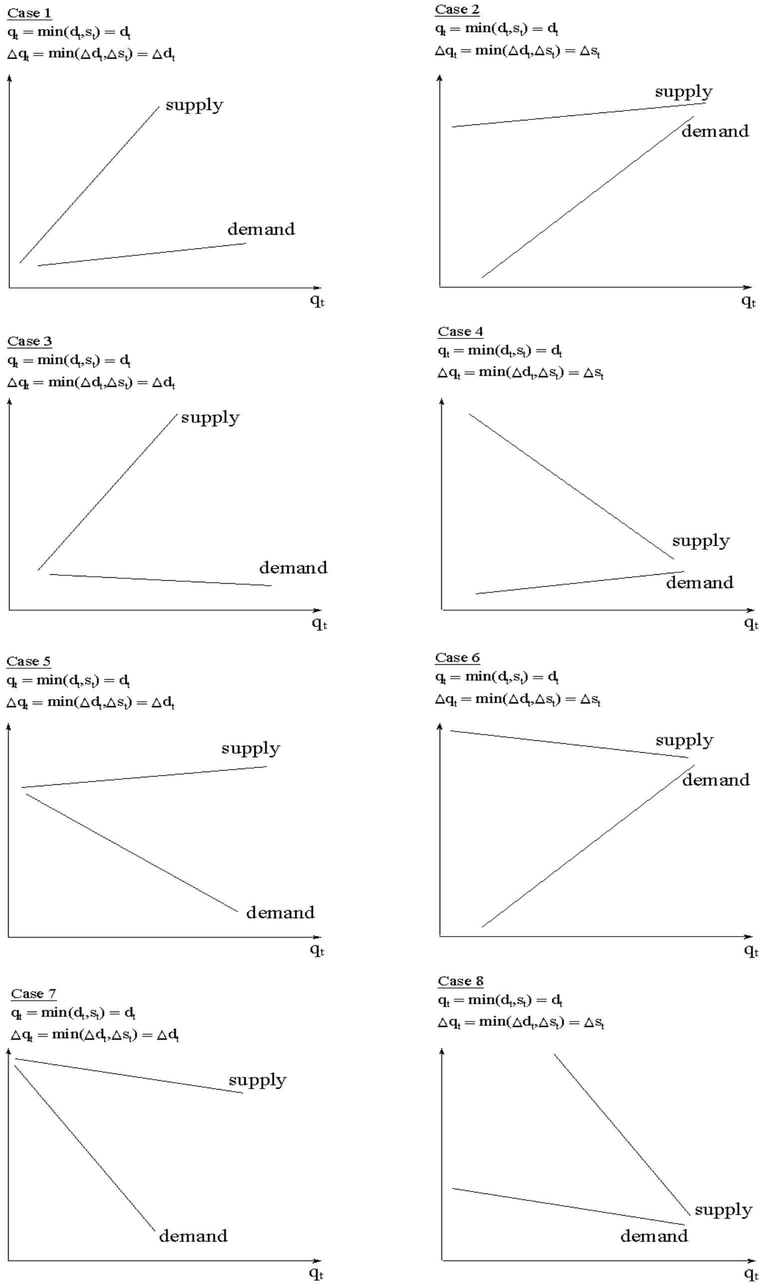
The above model allows for two regimes for characterizing Δq_t. Given equation (24), the growth rate of the amount of loans exchanged in the market corresponds to the minimum of the loan supply and demand growth rates. In other words, we consider that a demand (supply) regime occurs if the growth rate of the quantity of loans is determined by the variables and their parameters associated with the annual increase in loan demand (supply).

What are the advantages and drawbacks of this model? Firstly, it is obvious that, as a transition rule, the "min" condition defined on stationary data is not equivalent to the "min" condition defined on the level of the corresponding variables. The true DGP of Δq_t (if the standard disequilibrium model is valid) does not coincide with the model defined by the equations (22), (23) and (24). However (and even if it is not its primary objective), our model can be considered an approximation of the true DGP if the standard disequilibrium model is valid. For instance, let us consider the relationship between the type of disequilibrium in levels and in growth rates represented in Figure (2). In all cases contained therein, we assume that the true DGP of the observable quantity of loans, q_t, is determined by the level of demand (eight alternative and symmetric cases are also possible if we consider a supply regime). With these eight configurations, we explicitly take into account all the cases of positive and negative changes in supply and demand and their corresponding relative growth rates.

Two different groups of configurations can be distinguished. In the first group, there are four cases (numbered 1, 3, 5 and 7) in which the traditional disequilibrium model and our "min" condition give the same qualitative results, i.e. the identification of a demand regime. The growth rate of q_t is then equal to the growth rate of demand. On the contrary, in the second group, there are four cases (numbered 2, 4, 6 and 8) in which our "min" condition leads to a "wrong" regime identification, i.e. in which the growth rate of q_t is determined by Δs_t. Hence, in these cases, if the traditional disequilibrium model is the true DGP, the "min" condition on the growth rates leads to a wrong regime identification. For instance[11], it occurs when the (positive) changes in demand are very large and the (positive) changes in supply are smaller, whereas the market exhibits excess supply (as in case 2). Nevertheless, as can be observed in cases 2, 4, 6 and 8, these configurations are about to revert. In particular, they imply that the difference between loan supply and demand is decreasing. Hence, in all these cases, after a given delay, the demand will exceed the supply in levels and, as a consequence, our "min" condition will give a conclusion compatible with the true DGP. The level of q_t will be determined by s_t (excess demand) with the minimum of the growth rates still being equal to the growth rate of s_t. Yet even if these situations of wrong regime identification are transitory, they may last for a more or less protracted period of time.

[11] We are grateful to an anonymous referee for making this point.

Figure 2: Relationship between the type of disequilibrium in levels and in growth rates (the case of a demand regime)



Secondly, our model allows to (*i*) identify the disequilibrium in the loan market and (*ii*) to take into account the non-stationarity of the data. Indeed, the disequilibrium between demand and supply ($d_t \neq s_t$) implies an inequality of the corresponding growth rates ($\Delta d_t \neq \Delta s_t$). In other words, if there is a disequilibrium in level (in the traditional sense), it necessarily implies a divergence in the growth rates of the supplied and demanded quantities of loans. However, since the quantity of loans and thereby the estimates of the unobservable loan demand and supply variables are defined here as annual growth rates, our methodology precludes the identification of the type of disequilibrium in level form (whether the level of loan demand exceeds the level of loan supply or vice versa). For that reason, the aim of our approach is not to define the disequilibrium as an excess of supply or an excess of demand. The logic is different: given the non-stationarity issue, we propose to identify the disequilibrium as an excess of quantities of new loans supplied (or demanded) on the market at time t. For instance, if supply conditions deteriorate and the growth of supply is lower than the growth of demand, changes in the observed quantities of loans are assumed to be equal to the less restrictive growth rate, i.e. the growth of the supplied quantities, irrespective of the level of loan supply and demand. We are aware that the use of this "min" condition as a transition rule on the annual growth rates lacks theoretical foundations, based on the voluntary exchange principle. Yet, in this context, the disequilibrium model given by equations (22), (23) and (24) should simply be considered as a regime-switching model among others with a contemporaneous threshold variable. Moreover, this method provides an identification of the disequilibrium which is closely in line with the economic history of the Polish loan market. As a result, our main conclusion is that the disequilibrium has to be identified with stationary components of the observable data used. In our view, the abnormal results presented in the Section 3 stem from the misspecification of the system. It is possible to use several approaches in order to circumvent this problem. The solution that we propose has its advantages (it is technically easy to implement), but also its drawbacks (the method does not enable us to identify the type of disequilibrium considered in level form and we have only an approximation of the dynamic of Δq_t if the disequilibrium model $q_t = \min(d_t, s_t)$ is the true DGP).

Based on equations (22), (23) and (24), the estimation results for all empirical models are reported in Table (3).

——— Insert Table (3) ———

Starting with Model 1, the coefficients of the total deposits growth rate (DEP_t)

and of the proxy for annual increases in firms' cash flow (SAV_t) appear significant and have the expected signs. Moreover, the adjusted R-squared statistic is not as high as in the level form specification of the model (0.41 versus 0.991) and, more importantly, the frequency of the supply regime amounts to only 34 per cent as compared with 82 per cent obtained previously. In Model 2, the result of introducing the interest rates is mixed: they are significant and have correctly-signed coefficients only in the demand equation. Yet, the other variables are still significant, in particular the industrial production growth rate ($PROD_t$), which has a positively signed coefficient. Finally, although the quality of the model increased in comparison with Model 1, it is nevertheless still fairly low in terms of the $\bar{R}^2$ statistic. The quality of the model improves when changes in the imperfect substitutability between loans and risk-free assets (ATB_t) on the supply side are taken into account, and when changes in other forms of credit (OL_t) are integrated on the demand side. Indeed, in Model 3 there is a significant increase in the values of the log-likelihood and in the adjusted R-squared statistics. In addition, all the estimated parameters in both functions have the expected signs and, except for the interest rates in the loan supply function, are significant. As shown by Model 4 estimates, these findings are robust to the introduction of two extra variables reflecting some open economy aspects, i.e. annual increases in import prices (IMP_t) and changes in the net liquidity impact of foreign reserve accumulation by the central bank ($LNFA_t$). As for the level form estimates, we checked in Model 5 that our results are not affected by the existing collinearity between the interest rates. Finally, with a one-lag structure on the interest rates (cf. Model 6), all coefficients have the expected signs and are statistically significant.

Figure (3) plots the estimated probabilities of demand and supply regimes derived from Model 6 and Figure (4) plots the corresponding supply and demand growth rates. It should be recalled that the growth rate of the quantity of loans exchanged in the market corresponds to the minimum of the loan supply and demand growth rates. In other words, a demand (supply) regime occurs if the growth rate of loans is determined by the variables and their parameters associated with the annual increase in loan demand (supply). The results clearly reveal the existence of two regimes and corroborate with what one could expect on the Polish credit market: the protracted period of a demand regime observed until mid-1999 and the unambiguous apparition of a supply regime since then. Both regimes are in line with a number of stylized facts. To demonstrate this point, we use several external indicators, not included in the model. We also refer to different stylized facts about the economic history of the Polish loan market and to the information collected from the banks by the central bank.

Figure 3: Estimated Probabilities of Each Regime, III/94-II/02

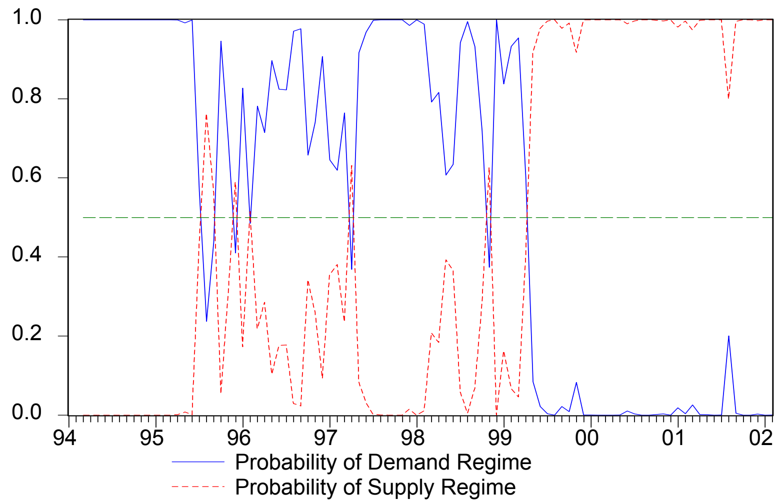
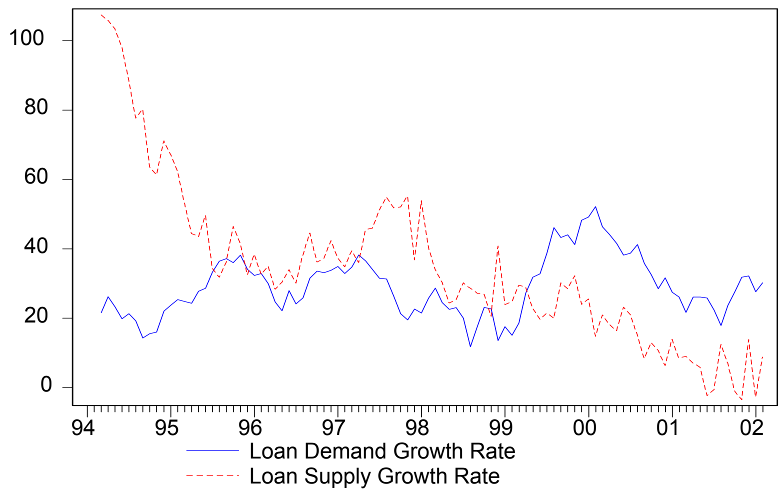


Figure 4: Estimated Demand and Supply Annual Growth Rates, III/94-II/02



The period under consideration begins with an outstanding, almost 110 per cent annual growth rate of loan supply. This result can be explained by the eagerness of banks to build up again their market share on a healthy basis following the government recapitalization program. This is especially the case given that the ratio of credit to the corporate sector to GDP was historically extremely low in 1994 (see Table (2)). Since then, the annual increase in loan supply steadily declined to approximately 44 per cent in April 1995, with a growth rate differential between loan supply and demand amounting to 20 per cent. Subsequently, despite a significant increase in loan demand growth between April and November 1995, the loan supply growth rate remained robust. The rise in the annual change in loan demand was concomitant with the strong pick-up in economic activity, as annual GDP growth accelerated from 5.2 per cent in 1994 to 7.0 per cent in 1995 (see Table (4)). At the same time, as Table (4) indicates, there were huge capital inflows under the fixed exchange rate regime, which led to a spectacular 150 per cent increase in the foreign reserves of the central bank in 1995. Yet, since these capital inflows were in fact imperfectly sterilized, they sustained the loan supply growth rate at around 40 per cent from April to December 1995.

More generally, the huge increases in gross official reserves of the central bank, at an average 52 per cent growth rate over the 1994-1998 period, coupled with the sterilized interventions by the monetary authorities had far reaching consequences on the supply side of the market. They led to a structural excess liquidity in the banking system, defined as a net indebtedness of the central bank towards commercial banks. The consequence for the transmission mechanism is that banks became less reactive to policy measures. Indeed, the policy rate, i.e., the rate of open market operations, instead of determining the marginal cost of banks' liabilities, started to play the role of a marginal investment rate.

——— Insert Table (4) ———

Increasing excess liquidity of banks along with a constant improvement of their loan portfolio as depicted by the decline in non-performing loans at an average annual pace of 23 per cent between 1994 and 1997 (see Table (4)) had a protracted impact on the loan supply growth rate. A supply-led boom occurred in the credit market from mid-1996. According to our estimates, the annual growth rate of loan supply climbed steeply from 30 per cent in July 1996 and reached an all time high (55 per cent) in November 1997, with the corresponding figures for the loan demand growth rate of 24 and 19.5 per cent respectively. The reaction of the central bank corroborates the view that the credit expansion was mainly led by supply factors and not demand-driven. As of mid-1996, in order to curb the credit activity of banks, the central bank started to progressively tighten its monetary policy by rising its intervention rate and the rate of required reserves. However, given that these measures did not have the expected impact on banks' behavior, thereby acknowledging the failure of traditional monetary instruments, the central bank resorted to an unprecedented measure. In mid-September 1997 it began to accept six and nine-month deposits directly from the public at above-market rates. This measure was designed to have two effects: reduce the amount of liquidity in the banking system and sharpen competition for funds in order to force banks to substantially increase their deposit rates and, subsequently, their lending rates (Szpunar, 1998).

This exceptional policy measure appears to have had a strong impact on the supply side of the market. Indeed, as of end-1997, the loan supply growth rate declined steeply and the growth rate differential between loan supply and demand was progressively reduced to almost zero in October and November 1998. The market switched to a supply regime in May 1999. Moreover, the discrepancy between loan supply and loan demand growth rates increased markedly as from December 1999. Since February 2000, the loan market was characterized by falling supply and demand growth rates. Figure (5) plots the intervention rate of the central bank, the differential in annual changes in the loan market (growth rate of supply − growth rate of demand), and both the annual growth rate of industrial production and of the ratio of non-performing loans to total loans for the corporate sector since the end of 1994[12]. Table (4) and Figure (5) provide some clues for the switch and the persistence of a supply regime in the loan market. First, there was an increase in the instability of real activity. Whereas the average annual growth rate of GDP was nearly 6 per cent in the 1994-1998 period, it decreased by over 30 per cent, to 4.1 per cent in 1999 and 4.0 per cent in 2000. It fell to only 1 per cent in 2001, reaching an all-time low of 0.2 per cent in the fourth quarter. Second, the supply conditions deteriorated tremendously at the end of the

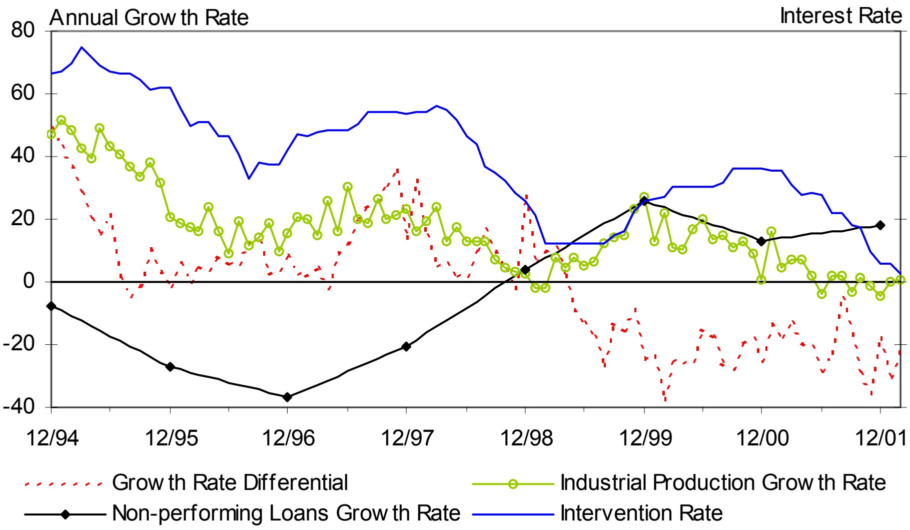
[12] We used the ratio for the entire banking sector only because the corresponding data for our loan series and for our sample of banks were not available.

1990s and the early 2000s. Non-performing loans started to rise six months before the switch to the supply regime and their growth rate skyrocketed in 1999. Changes in the gross profit to income ratio became negative in 1997 and recorded an over 50 per cent decline in 1998 as compared to the previous year. Both indicators continued to worsen in the subsequent years. External information collected by the NBP is line with our analysis. On-site examinations at banks pointed out that deteriorating economic conditions and higher default risk have led banks to apply stricter lending criteria (NBP, 2003a). Moreover, corporate surveys revealed an increase in the frequency of loan applications being refused by the banks from 12.8% in 1999 to 33.9% in 2002 (NBP, 2003b). The frequency of loan applications rejected for the first time and sporadically more than doubled, those refused frequently - more than tripled amounting to 13.5 per cent of surveyed companies seeking loans. In sum, according to the Polish central bank assessment, banks' behavior was the main driving force behind the declining expansion of bank lending (NBP, 2003a). Cross-border credit flows data is another indicator which tends to corroborate the finding of a supply regime in the Polish loan market since the end of the nineties. Indeed, as evidenced by balance of payments statistics, there was a huge (63.3 per cent) increase in cross-border credit flows between 1999 and 2002. Consequently, it appears that foreign banks were taking up the slack, at least in part, of a weaker supply than demand growth rate in the domestic loan market[13].

Finally, it is difficult to establish exactly what kind of shock brought about a regime change in the loan market. This issue must be analyzed with caution. Nevertheless, as Figure (5) indicates, the quasi-simultaneity of a supply regime, monetary tightening, the waning quality of extended loans and the decline in the industrial production growth rate is striking. At any rate, the coexistence of a rising monetary policy rate and of a supply regime means that banks acted pro-cyclically in the business cycle, thus amplifying the consequences of an increasingly restrictive monetary policy.

[13] We are grateful to an anonymous referee for making this point.

Figure 5: Growth Rate Differential Between Loan Supply and Demand and Economic Conditions, XII/94-II/02



6 Conclusion

The Polish credit market conditions changed dramatically during the 1990s and the clearing of the market was not completely secured by the adjustments of interest rates. The objective of our paper is to confirm econometrically the switch observed in the Polish loan market and to test the disequilibrium hypothesis. For this reason, we apply the standard ML method on a disequilibrium model with variables in level and a short-side rule. However, the results are very counter-intuitive, in particular as regards the identification of the regimes in the February 1994 - February 2002 period. It seems that, in our case, the implementation of this technique does not allow us to identify the type of disequilibrium *a priori* expected on the Polish loan market.

We propose an explanation of these counter-intuitive results based on the non-stationarity of the data used. We show that almost all observed variables are derived from non-stationary processes and are not cointegrated in the corresponding linear models. The issue in the disequilibrium model is that it is impossible to test the cointegration hypothesis in demand and supply equations since demand and supply quantities are unobservable. Besides, if residuals in demand and supply equations are non-stationary (*spurious regression representation*), the incorrect use of the standard log-likelihood function (Maddala and Nelson, 1974) may lead to an erroneous identification of demand and supply regimes.

In order to asses this proposition, we estimate the disequilibrium model with stationary data in first differences. In a first approximation of the transition rule, we use the "min" condition on the annual growth rates. This approximation does not reflect the exact dynamic of the growth rate of the observed loan quantities. However, it leads to a very accurate identification of regimes on the Polish credit market in the 1990s. Hence, the determination of the appropriate ML estimation method in the *spurious regression representation* is the next goal in our research program.

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A Appendix

A.1 Marginal densities of Q_t in a stable disequilibrium model

Let us denote $g_{D_t,S_t}(d_t, s_t)$ the joint density of D_t and S_t. We know that the corresponding marginal densities of the unobservable variables D_t and S_t are defined by:

$$f_{D_t}(d_t) = \int_{-\infty}^{\infty} g_{D_t,S_t}(d_t, z)\, dz \qquad f_{S_t}(s_t) = \int_{-\infty}^{\infty} g_{D_t,S_t}(z, s_t)\, dz \tag{25}$$

We have to compute the marginal density of Q_t on the two subset $Q_t = D_t$, with $D_t < S_t$ and $Q_t = S_t$, with $S_t < D_t$. When $D_t < S_t$, for a given realization d_t of D_t, the marginal density of Q_t, is given by the area defined by the joint density $g_{D_t,S_t}(d_t, z)$, for values z of S_t superior to d_t. Under the assumption that $D_t < S_t$, the marginal density of Q_t is then given by:

$$f_{Q_t|D_t<S_t}(q_t) = \int_{q_t=d_t}^{\infty} g_{D_t,S_t}(d_t, z)\, dz \tag{26}$$

Symmetrically, we obtain the marginal density of Q_t when $S_t < D_t$.

$$f_{Q_t|S_t<D_t}(q_t) = \int_{q_t=s_t}^{\infty} g_{D_t,S_t}(z, s_t)\, dz \tag{27}$$

In the general case, we know that the marginal density of Q_t is given by:

$$f_{Q_t}(q_t) = \int_{q_t}^{\infty} g_{D_t,S_t}(q_t, z)\, dz + \int_{q_t}^{\infty} g_{D_t,S_t}(z, q_t)\, dz \tag{28}$$

Let us assume that both residuals ε_1 and ε_2 are independent ($\sigma_{12} = 0$). In this case, the joint density can be expressed as the following simple expression:

$$\begin{aligned} g_{D_t,S_t}(d_t, s_t) &= \frac{1}{2\pi\sigma_1\sigma_2} \exp\left\{-\frac{1}{2}\left[\left(\frac{d_t - x_{1,t}'\beta_1}{\sigma_1}\right)^2 + \left(\frac{s_t - x_{2,t}'\beta_2}{\sigma_2}\right)^2\right]\right\} \\ &= \frac{1}{2\pi\sigma_1\sigma_2} \exp\left[-\frac{1}{2}\left(\frac{d_t - x_{1,t}'\beta_1}{\sigma_1}\right)^2\right] \times \exp\left[-\frac{1}{2}\left(\frac{s_t - x_{2,t}'\beta_2}{\sigma_2}\right)^2\right] \end{aligned} \tag{29}$$

Now, consider the first member of the marginal density of Q_t (equation (28)):

$$\int_{q_t}^{\infty} g_{D_t,S_t}(q_t, z)\, dz = \frac{1}{2\pi\sigma_1\sigma_2} \int_{q_t}^{\infty} \left\{\exp\left[-\frac{1}{2}\left(\frac{q_t - x_{1,t}'\beta_1}{\sigma_1}\right)^2\right] \times \exp\left[-\frac{1}{2}\left(\frac{z - x_{2,t}'\beta_2}{\sigma_2}\right)^2\right]\right\} dz$$

$$= \frac{1}{\sqrt{2\pi}\sigma_1} \exp\left[-\frac{1}{2}\left(\frac{q_t - x_{1,t}'\beta_1}{\sigma_1}\right)^2\right] \times \frac{1}{\sqrt{2\pi}\sigma_2} \int_{q_t}^{\infty} \exp\left[-\frac{1}{2}\left(\frac{z - x_{2,t}'\beta_2}{\sigma_2}\right)^2\right] dz$$

In the first term of this expression, we recognize the value of the $N(0,1)$ density function at the particular point $\left(q_t - x'_{1,t}\beta_1\right)/\sigma_1$. Indeed:

$$\frac{1}{\sqrt{2\pi}} \exp\left[-\frac{1}{2}\left(\frac{q_t - x'_{1,t}\beta_1}{\sigma_1}\right)^2\right] = \phi\left(\frac{q_t - x'_{1,t}\beta_1}{\sigma_1}\right) \tag{30}$$

where $\phi(.)$ denotes the $N(0,1)$ density function. Since this function is symmetric the first member of the marginal density of Q_t can be expressed as:

$$\int_{q_t}^{\infty} g_{D_t,S_t}(q_t, z)\, dz = \frac{1}{\sigma_1}\phi\left(\frac{x'_{1,t}\beta_1 - q_t}{\sigma_1}\right) \times \frac{1}{\sqrt{2\pi}\sigma_2}\int_{q_t}^{\infty} \exp\left[-\frac{1}{2}\left(\frac{z - x'_{2,t}\beta_2}{\sigma_2}\right)^2\right] dz \tag{31}$$

The second term of this expression can be transformed in order to introduce the $N(0,1)$ cumulative distribution function, denoted $\Phi(.)$. Indeed, let us consider the following change in variable $\widetilde{z} = \left(z - x'_{2,t}\beta_2\right)/\sigma_2$, with $dz = d\widetilde{z}\,\sigma_2$. Then, we have:

$$\begin{aligned} \frac{1}{\sqrt{2\pi}\sigma_2}\int_{q_t}^{\infty} \exp\left[-\frac{1}{2}\left(\frac{z - x'_{2,t}\beta_2}{\sigma_2}\right)^2\right] dz &= \frac{1}{\sqrt{2\pi}\sigma_2}\int_{\widetilde{q}_t}^{\infty} \exp\left(-\frac{\widetilde{z}^2}{2}\right) d\widetilde{z}\,\sigma_2 \quad (32) \\ &= \frac{1}{\sqrt{2\pi}}\int_{\widetilde{q}_t}^{\infty} \exp\left(-\frac{\widetilde{z}^2}{2}\right) d\widetilde{z} \end{aligned}$$

with $\widetilde{q}_t = \left(q_t - x'_{2,t}\beta_2\right)/\sigma_2$. Then, this integral can be expressed as a function $\Phi(.)$.

$$\frac{1}{\sqrt{2\pi}\sigma_2}\int_{q_t}^{\infty} \exp\left[-\frac{1}{2}\left(\frac{z - x'_{2,t}\beta_2}{\sigma_2}\right)^2\right] dz = 1 - \Phi(\widetilde{q}_t) = \Phi(-\widetilde{q}_t) \tag{33}$$

Finally, we obtain:

$$\int_{q_t}^{\infty} g_{D_t,S_t}(q_t, z)\, dz = \frac{1}{\sigma_1}\phi\left(\frac{x'_{1,t}\beta_1 - q_t}{\sigma_1}\right)\Phi\left(\frac{x'_{2,t}\beta_2 - q_t}{\sigma_2}\right) \tag{34}$$

Symmetrically, we can compute the second term of the marginal density of Q_t (equation (28)) as:

$$\int_{q_t}^{\infty} g_{D_t,S_t}(z, q_t)\, dz = \frac{1}{\sigma_2}\phi\left(\frac{x'_{2,t}\beta_2 - q_t}{\sigma_2}\right)\Phi\left(\frac{x'_{1,t}\beta_1 - q_t}{\sigma_1}\right) \tag{35}$$

Then, the marginal density of Q_t is defined by equation (10):

$$\begin{aligned} f_{Q_t}(q_t) &= \frac{1}{\sigma_1}\phi\left(\frac{x'_{1,t}\beta_1 - q_t}{\sigma_1}\right)\Phi\left(\frac{x'_{2,t}\beta_2 - q_t}{\sigma_2}\right) \\ &\quad + \frac{1}{\sigma_2}\phi\left(\frac{x'_{2,t}\beta_2 - q_t}{\sigma_2}\right)\Phi\left(\frac{x'_{1,t}\beta_1 - q_t}{\sigma_1}\right) \end{aligned}$$

A.2 The choice of initial conditions in the *ML* optimization procedure

There are various methods to obtain the initial conditions on structural parameters θ in the ML iteration. Here, we use a two-step OLS procedure. First, we consider the linear regressions of the observation q_t on the exogenous variables sets in both functions: $q_t = x'_{i,t}\widehat{\gamma}_i + \mu_{i,t}$, with $i = 1, 2$. Given the realizations of $\widehat{\gamma}_1$ and $\widehat{\gamma}_2$, we compute a first approximation of demand and supply, as $\widetilde{d}_t = x'_{1,t}\widehat{\gamma}_1$ and $\widetilde{s}_t = x'_{2,t}\widehat{\gamma}_2$. Even if we know that $\widehat{\gamma}_1$ and $\widehat{\gamma}_2$ are not convergent estimators of β_1 and β_2, we build two subgroups of observations. In the first subgroup, denoted by index d, we consider only the observations on Q_t, $X_{1,t}$ and $X_{2,t}$ for which we have $\widetilde{d}_t \leq \widetilde{s}_t$. In the second subgroup, we consider the observations for which we have $\widetilde{s}_t \leq \widetilde{d}_t$. The second step of the procedure consists in applying the OLS on both subgroups:

$$q_t^{(d)} = x_{1,t}^{(d)\prime}\widetilde{\beta}_1 + \widetilde{\mu}_{1,t} \quad \text{and} \quad q_t^{(s)} = x_{2,t}^{(s)\prime}\widetilde{\beta}_2 + \widetilde{\mu}_{2,t} \tag{36}$$

Then, we use the OLS estimates $\widetilde{\beta}_i$ as starting values for β_i in the ML iteration. For the parameters σ_1 and σ_2, we adopt the following starting values:

$$\widetilde{\sigma}_i = \frac{1}{n_i}\sum_{j=1}^{n_i}\widetilde{\mu}_{i,j} \quad i = 1, 2 \tag{37}$$

where n_1 denotes the size of the "demand" subgroup of observations for which we have $\widetilde{d}_t \leq \widetilde{s}_t$, and n_2 denotes the size of the corresponding "supply" subgroup. Some Monte Carlo simulations of the accuracy of this procedure are available upon request.

A.3 Unit root and cointegration tests

The results of the unit root and cointegration tests are reported in Tables (5) and (6).

——— Insert Table (5) ———

——— Insert Table (6) ———

Table 1: ML Estimates in Level, II/94 - II/02

	Model 1	Model 2	Model 3	Model 4	Model 5	Model 6
Supply						
Constant	-0.21^{**} (-4.00)	-0.65^{**} (-3.24)	0.78^{*} (1.66)	0.32 (0.74)	0.32 (0.74)	0.49 (1.57)
IL_t		0.004^{**} (2.57)	$\mathbf{-0.003}$ (-1.51)	$\mathbf{-0.0003}$ (-0.13)		
IC_t		-0.002 (-1.48)	$\mathbf{0.003}^{*}$ (1.71)	-0.0002 (-0.11)	-0.0005 (-0.39)	
IL_{t-1}						$\mathbf{-0.0003}$ (-0.17)
IC_{t-1}						-0.001 (-1.07)
$(IL-IC)_t$					-0.0003 (-0.13)	
DEP_t	0.90^{**} (82.85)	0.98^{**} (27.62)	0.77^{**} (11.29)	0.77^{**} (11.81)	0.77^{**} (11.81)	0.73^{**} (15.91)
ATB_t			-0.27^{**} (-3.28)	-0.19^{**} (-2.56)	-0.19^{**} (-2.56)	-0.16^{**} (-3.01)
$LNFA_t$				0.17^{**} (3.47)	0.17^{**} (3.47)	0.17^{**} (4.06)
σ_1	0.021	0.019	0.019	0.016	0.016	0.015
Demand						
Constant	3.65^{**} (14.44)	3.46^{**} (7.44)	5.86^{**} (6.12)	2.88^{**} (3.05)	2.88^{**} (3.04)	5.40^{**} (6.82)
IL_t		-0.02^{**} (-5.98)	-0.001 (-0.31)	-0.006^{*} (-1.77)		
IC_t		0.02^{**} (6.11)	$\mathbf{-0.000}$ (-0.005)	0.003 (1.31)	-0.002 (-1.53)	
IL_{t-1}						-0.02^{**} (-8.35)
IC_{t-1}						0.01^{**} (7.95)
$(IL-IC)_t$					-0.006^{*} (-1.77)	
$PROD_t$		-0.09 (-0.71)	-0.02 (-0.26)	-0.03 (-0.27)	-0.03 (-0.27)	0.12^{**} (2.24)
SAV_t	$\mathbf{0.21}^{**}$ (3.71)	$\mathbf{0.39}^{**}$ (4.23)	$\mathbf{0.65}^{**}$ (9.88)	$\mathbf{0.51}^{**}$ (7.77)	$\mathbf{0.51}^{**}$ (7.75)	$\mathbf{0.28}^{**}$ (4.36)
OL_t			-2.20^{**} (-4.14)	-0.99^{**} (-2.24)	-0.99^{**} (-2.24)	-1.80^{**} (-5.20)
IMP_t				0.61^{**} (3.41)	0.61^{**} (3.40)	0.34^{**} (3.45)
σ_2	0.008	0.013	0.010	0.010	0.010	0.003
	$L = 254.77$	$L = 256.62$	$L = 271.20$	$L = 282.15$	$L = 282.15$	$L = 287.80$
	$\bar{R}^2 = 0.991$	$\bar{R}^2 = 0.992$	$\bar{R}^2 = 0.993$	$\bar{R}^2 = 0.994$	$\bar{R}^2 = 0.994$	$\bar{R}^2 = 0.994$
	$n_S = 77$	$n_S = 66$	$n_S = 66$	$n_S = 67$	$n_S = 67$	$n_S = 79$
	$n_D = 17$	$n_D = 26$	$n_D = 30$	$n_D = 30$	$n_D = 30$	$n_D = 16$
	$F_S = 82\%$	$F_S = 72\%$	$F_S = 69\%$	$F_S = 69\%$	$F_S = 69\%$	$F_S = 83\%$
	$F_D = 18\%$	$F_D = 28\%$	$F_D = 31\%$	$F_D = 31\%$	$F_D = 31\%$	$F_D = 17\%$

Source: Authors' calculations.

Notes: (1) Asymptotic t-statistics are in parentheses, L: log-likelihood. (2) * (**) denotes significance at 10% (5%) level. (3) Bold characters are for variables which do not have the expected sign coefficient. (4) nD (nS) denotes the number of demand (supply) periods, and FD (FS) denotes the frequency of supply (demand) regimes given that a period or regime occurs when the corresponding probability is higher than 0.5.

Table 2: Credit to the Corporate Sector to GDP Ratio, 1989 - 1999

1989	1990	1991	1992	1993	1994	1995	1996	1997	1998	1999
28.3	21.1	23	20.6	19.9	17.1	16.4	17.6	18.6	19.8	21.2

Source: National Bank of Poland and authors' calculations.

Table 3: ML Estimates in Growth Rates, II/94 - II/02

	Model 1	Model 2	Model 3	Model 4	Model 5	Model 6
ΔSupply						
Constant	-11.69 (-1.25)	7.85 (0.37)	-1.54 (-0.09)	-6.29 (-0.32)	-6.29 (-0.32)	-30.77^{**} (-2.46)
IL_t		$\mathbf{-3.09}$ (-0.95)	0.72 (0.30)	1.74 (0.51)		
IC_t		$\mathbf{1.92}$ (0.68)	-1.80 (-0.96)	-2.99 (-1.01)	-1.25 (-1.19)	
IL_{t-1}						7.10^{**} (2.83)
IC_{t-1}						-7.20^{**} (-3.16)
$(IL-IC)_t$					1.73 (0.51)	
ΔDEP_t	1.63^{**} (2.95)	2.12^{**} (2.45)	1.74^{**} (4.03)	1.94^{**} (4.03)	1.94^{**} (4.03)	1.37^{**} (7.10)
ΔATB_t			-0.44^{**} (-3.67)	-0.44^{**} (-2.95)	-0.44^{**} (-2.95)	-0.56^{**} (-5.26)
$\Delta LNFA_t$				0.31^{**} (3.40)	0.31^{**} (3.40)	0.33^{**} (4.21)
σ_1	10.62	10.42	7.57	7.05	7.05	5.87
ΔDemand						
Constant	37.13^{**} (13.38)	40.55^{**} (4.42)	38.97^{**} (6.04)	33.56^{**} (7.64)	33.56^{**} (7.64)	49.87^{**} (10.07)
IL_t		-1.32^{**} (-2.64)	-1.40^{**} (-4.96)	-1.40^{**} (-4.41)		
IC_t		1.07 (1.52)	1.78^{**} (4.06)	1.73^{**} (4.16)	0.32 (1.27)	
IL_{t-1}						-1.38^{**} (-4.24)
IC_{t-1}						0.92^{**} (2.07)
$(IL-IC)_t$					-1.40^{**} (-4.41)	
$\Delta PROD_t$		0.34^{**} (2.53)	0.04 (0.37)	-0.04 (-0.50)	-0.04 (-0.50)	0.07 (0.89)
ΔSAV_t	-0.29^{**} (-4.17)	-0.33^{**} (-4.31)	-0.27^{**} (-6.49)	-0.26^{**} (-6.95)	-0.26^{**} (-6.95)	-0.25^{**} (-5.90)
ΔOL_t			-2.97^{**} (-7.45)	-1.77^{**} (-5.45)	-1.77^{**} (-5.45)	-2.46^{**} (-6.75)
ΔIMP_t				0.27^{**} (2.85)	0.27^{**} (2.85)	0.31^{**} (3.47)
σ_2	5.71	4.33	2.72	3.05	3.05	2.74
	L=−324.84	L=−308.10	L=−284.95	L=−272.04	L=−272.04	L=−261.72
	$\bar{R}^2 = 0.41$	$\bar{R}^2 = 0.47$	$\bar{R}^2 = 0.63$	$\bar{R}^2 = 0.73$	$\bar{R}^2 = 0.73$	$\bar{R}^2 = 0.80$
	$n_S = 31$	$n_S = 30$	$n_S = 53$	$n_S = 32$	$n_S = 32$	$n_S = 38$
	$n_D = 61$	$n_D = 66$	$n_D = 43$	$n_D = 63$	$n_D = 63$	$n_D = 56$
	$F_S = 34\%$	$F_S = 31\%$	$F_S = 55\%$	$F_S = 34\%$	$F_S = 34\%$	$F_S = 40\%$
	$F_D = 66\%$	$F_D = 69\%$	$F_D = 45\%$	$F_D = 66\%$	$F_D = 66\%$	$F_D = 60\%$

Source: Authors' calculations.

Notes: (1) Asymptotic t-statistics are in parentheses, L: log-likelihood. (2) * (**) denotes significance at 10% (5%) level. (3) Bold characters are for variables which do not have the expected sign coefficient. (4) nD (nS) denotes the number of demand (supply) periods, and FD (FS) denotes the frequency of supply (demand) regimes given that a period or regime occurs when the corresponding probability is higher than 0.5.

Table 4: GDP, Gross Official Reserves and Banking Indicators, 1994 - 2001 (Annual Growth Rates, in per cent)

	1994	1995	1996	1997	1998	1999	2000	2001
GDP	5.2	7.0	6.0	6.8	4.8	4.1	4.0	1.0
Gross official reserves (end of year)	39.5	150.0	20.0	15.0	36.7	-3.5	0.7	-3.3
Non-performing loans (end of year)	-7.4	-27.2	-36.8	-20.5	3.8	25.7	13.1	18.1
Gross profit/income (end of year)	0.0	181.7	14.8	-24.2	-51.7	-12.7	-25.8	-43.5

Source: OECD (2000), IMF(2002), and National Bank of Poland.

Table 5: Augmented Dickey-Fuller and Phillips-Perron Tests: Variables in Level Forms, II/94 - II/02

	ADF Statistics						PP Statistics	
Variable	Numb. of Lags	UR with Const.	Joint Test: UR and no Const.	T-Statistic on the Const.	Numb. of Lags	UR without Const.	UR with Const.	UR without Const.
IL	1	-1.93	4.33*	-	1	-2.63**	-1.89	-3.05**
IC	6	-1.41	1.56	-	1	-1.84**	-0.39	-1.73*
CR	0	-2.37	24.38**	-	2	3.11	-2.22	5.40
$PROD$	8	-3.15**	-	3.22**	8	1.96	-2.93**	2.24
DEP	2	-5.34**	-	5.79**	1	7.22	-5.27**	9.58
SAV	0	-1.40	6.77**	-	2	4.47	-1.63	4.51
$LNFA$	0	-2.48	3.36	-	1	0.74	-2.16	0.94
IMP	0	-2.67*	5.41**	-	0	1.77	-3.03**	2.11
OL	0	-1.48	1.71	-	0	1.08	-1.48	1.10
ATB	0	-0.63	0.57	-	0	-0.91	-0.61	-0.94

Source: Authors' calculations.

Notes: (1) * (**) denotes rejection of the null hypothesis at 10% (5%) level. (2) The optimal lag number for ADF tests was determined using the Schwarz criterion, and for PP tests according to the Newey-West criterion, which indicated 3 lags.

Table 6: Johansen Cointegration Tests Between the Amount of Credit and Exogenous Variables, II/94 - II/02

	Model 1	Model 2	Model 3	Model 4	Model 5	Model 6
H_0	$R = 0$	$R = 0$	$R = 0$	$R = 0$	$R = 0$	$R = 0$
trace statistics	2.09	1.30	18.15**	32.13**	32.13**	35.68**

Source: Authors' calculations.

Notes: (1) ** denotes rejection of the null hypothesis at 5% significance level. (2) Tests allow for an intercept but no trend in the cointegrating relationship.